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Structural Breaks and Long-Run Trends in Commodity Prices

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Time-series techniques that control for the presence of structural breaks reveal that the international price of most commodities presents a negative long-run trend for 1900–92. They also show that shocks are far less persistent than previously estimated. Both findings suggest that there may be important welfare gains from stabilization mechanisms such as commodity funds.

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Summary findings

The oil shocks of the 1970s, which quadrupled the price of petroleum, marked the end of an abnormal period of price stability and renewed interest in predicting the evolution of commodity prices. But most subsequent studies have focused on the short-run effects of price fluctuations, mainly because they greatly affect the foreign trade of developing countries. Sophisticated compensation mechanisms, such as commodity funds, have been introduced to counterbalance the transitory effects of price shocks.

But the long-term evolution of prices also affects policy design and development strategies and may have a more important role in fostering long-run growth. The evidence presented by Prebisch (1950) and Singer (1950) of a secular negative trend in the price of commodities in 1870–1945 implies an increasingly weak position for developing countries relative to industrial economies. This hypothesis by Prebisch and Singer has been strongly debated, both theoretically and empirically, during the past four decades.

Using recent advances in econometric theory, León and Soto analyze the long-run dynamics of the price of the 24 most-traded commodities in 1900–92. The method they use tests for nonstationarity (unit roots) in the series with a technique that allows structural breaks to be endogenously determined. The results show that 15 of the 24 commodity prices present negative trends, six are trendless, and three exhibit positive trends. Thus, the Prebisch-Singer hypothesis, though not universal, holds for most commodities. This evidence rejects, to some extent, previous evidence by Cuddington (1992) and others.

León and Soto extend the econometric analysis to determine the persistence of shocks to commodity prices. Knowledge of the persistence of shocks is important when designing counterbalancing policies such as commodity funds. The authors use a nonparametric estimator of persistence (the multiple variance ratio) and find that 19 of the 24 commodity prices present persistence levels substantially lower than previous estimates. This evidence suggests that there may be substantial room for stabilization and price support mechanisms for most commodities.

This paper — a product of the Macroeconomics and Growth Division, Policy Research Department — is part of a larger effort in the department to understand the links of foreign shocks and macroeconomic policies. Copies of the paper are available free from the World Bank, 1818 H Street NW, Washington, DC 20433. Please contact Raquel Luz, room N11-043, extension 31320 (24 pages). January 1995.

STRUCTURAL BREAKS AND LONG-RUN TRENDS IN COMMODITY PRICES

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1. Introduction.

The oil shock, which increased the price of petroleum by four times twenty years ago, not only marked the end of a period of price stability totally abnormal in economic history, but also renewed the interest of economists on the short and long run evolution of commodity prices. Most of the subsequent theoretical and empirical studies, however, have concentrated on the short-run effects of price fluctuations rather than in their long-run swings, mainly because of the dependence of developing countries on foreign trade. The harmful effects of short-term commodity price fluctuations on the performance of small economies, extensively documented in the Dutch disease literature, has forced LDCs to introduce sophisticated compensation mechanisms, aimed at counterbalancing the transitory effects of terms of trade shocks.

Despite the importance of short-run fluctuations, the long-term evolution of prices also plays a crucial role for policy design and development strategies, since fluctuating export revenues affect exchange rate management, fiscal revenues and, ultimately, long-run growth as exemplified by the supply-side shocks of the 1970s. This issue has been a matter of economic interest probably since Adam Smith's Wealth of Nations, although during the last quarter of a century the topic has received less attention from academics than from policy-makers. The most influential studies have been the contributions of Prebisch (1950) and Singer (1950), which present theoretical justification and empirical evidence of a secular negative trend in the price of primary commodities relative to that of manufactured goods in the 1870-1945 period, a postulate known as the Prebisch-Singer hypothesis (PSH). The nature of this decline --which favors manufactures-exporting countries to the detriment of less developed primary goods producers-- has been justified by a combination of productivity differentials in favor of the former, asymmetric market structures (oligopolic rents in manufactures versus zero profit competition in primary goods) and higher income-elasticity in industrial than primary goods.[1]

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[1] An important implication of the PSH, which is at the core of structuralist theory, is that a secular decline in the terms of trade implies a progressively unfavorable international trade scenario for developing countries. The import-substitution strategy, undertaken during the 1950-1975 period in most Latin American countries, proposed a transitory withdrawal of the countries from world markets in order to develop a domestic industrial base, with which the economy would re-enter world markets in a stronger position (see Lustig, 1988).

Recent empirical evidence, however, has rejected the PSH using stochastic-trend models to overcome some technical limitations of early studies, which did not take into account the potential non-stationarity of commodity prices. Cuddington and Urzúa (1989) showed that the Grilli-Yang index of commodity prices should be represented as an integrated variable, and that the corresponding first-difference model did not contain a significant long-run trend in the 1900-1983 period. Cuddington (1992) applied this methodology to the analysis of the 24 commodities that comprise the index (plus oil and coal), and found that 21 of them present zero or positive trends, thus rejecting the PSH in the majority of the cases.[2]

Apart from the rejection of the PSH, an additional contribution of these papers is the analysis of the persistence of shocks. Knowledge of the persistence of shocks is important for the design of short and long-term policies. For example, the sequencing of structural adjustment programs with depressed terms of trade is certainly very different if export prices are expected to recover quickly or to remain depressed for a long period. Likewise, much of the new literature on commodity stabilization funds has emphasized that the optimal management of these funds depends to an important degree on the nature of the shocks to commodity prices, in particular their persistence and the speed at which they dissipate (Arrau and Claessens, 1991; Engel and Meller, 1993).

Nevertheless, this stochastic-trends approach presents two important limitations, that might have misguided the analysis of the PSH and the persistence of commodity price shocks. First, it relies crucially on an appropriate characterization of commodity prices as trend or difference stationarity processes (henceforth, TS and DS respectively). Widely used methods for verifying non-stationarity, in particular parametric tests such as the Dickey-Fuller test, are ill-suited for long-run analysis not only because they usually have low power against close competing alternatives, but also because they are particularly sensitive to structural breaks, likely to be present in long horizons. The second drawback, as noted by Cochrane (1991), is that even in the case that a unit-root test could successfully distinguish between competing TS and DS

[2] The empirical evidence for *aggregate* commodity prices is mixed; while Grilli and Yang (1988), Armani and Wright (1993) and Reinhart and Wickham (1994) found evidence of declining long-run trends using different datasets and time periods, Spraos (1980) and Cuddington and Urzua (1989) rejected the PSH.

models, the classification of a series may be of little help for econometric purposes if, as usual, the estimation procedure is performed in small samples.[3]

This paper revises the long and short run time-series structure of 24 commodity prices in order to answer the questions of the secular decline and the long-run persistence of shocks. Our methodology inverts, to some extent, the procedure used by Cuddington and Urzúa, which first tests for unit-roots, then models prices as TS or DS processes and, finally, infers the persistence of shocks (defined as the part of a shocks that have permanent effects on the level of commodity prices) by using the estimated gain functions of the ARIMA representation of DS models (shocks in TS models have zero persistence). The alternative procedure used in this paper evaluates first the persistence of shocks, using a non-parametric methodology which avoids most of the limitations of parametric tests and, in a second stage, tests long-run trends according to a model which can be parameterized better in small samples and taking into account the presence of structural breaks.

The non-parametric estimator of persistence used in this paper was proposed by Cochrane (1988 and 1991) and extended by Lo and McKinley (1989) and Chow and Denning (1993). It measures the importance of the non-stationary component of a series by comparing the variance of the permanent component to that of the innovations. Its main advantages are that it captures long-run mean-reversions more accurately than standard parametric techniques in small samples and, as shown below, is less sensitive to structural breaks and other econometric problems that plague parametric unit-root tests. Determining the persistence of shocks and the length of time required for the variable to revert to its mean growth rate allows a more appropriate representation of commodity prices.

The remainder of the paper is organized as follows. A discussion of the methodological limitations of early studies of the PSH is undertaken in section 2. Section 3 evaluates the loss of power of parametric unit-root test in the presence of structural breaks and determines the

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[3] The limited information contained in a small sample makes it hard to parameterize the true data generating process of a variable. If a TS model presents long-memory shocks, a small sample may not contain enough information to assess the long-run persistence of the shocks. In this case the variable resembles a non-stationary process, so that a DS specification would be more appropriate. On the other hand, if the variable is a combination of transitory and long-memory shocks, and the latter are relatively small with regards to the former, the series will be hardly distinguishable from a TS representation (see Cochrane, 1988).

existence of breaks in the commodity price data using an endogenous-break test. The non-parametric estimator of the long-run persistence of shocks to commodity prices is presented in section 4, as well as an evaluation of its performance under structural breaks. Section 5 undertakes an empirical reassessment of the Prebisch-Singer hypothesis and a taxonomy of the persistence and dynamics of commodity price shocks. Section 6 presents the conclusions and discusses some policy implications of the results.

2. Testing the Prebisch-Singer Hypothesis: A Methodological Review.

Several studies have tried to verify empirically the PSH but a clear answer remains elusive. Prebisch (1950) and Lewis (1952) provide early estimates of a negative trend in the terms of trade of LDCs in the 1870-1945 period, in the range of 0.5% to 0.7% per year. These estimates may appear small on a year-to-year basis, but can build up an important cumulative effect; in a 70 year period the terms of trade of LDCs would have declined by 35% to 45%. An initial refutation of the PSH is provided by Spraos (1980), who found a negative trend for the 70-year period ending in the outbreak of World War II, but non-significant results when the 1945-1970 period was added.

An important limitation of these and other early studies (eg, Sapsford, 1985) was the deficient quality of the data on terms of trade, which led the researchers to concentrate on the evolution of the commodity prices which comprise the term of trade indices. Grilli and Yang (1988) undertook a major effort to generate a reliable database for the 1900-1986 period, with which they tested for long-run trends under alternative definitions of commodity price indices and deflators. They found a negative trend for the entire period, similar in magnitude to those estimated by Prebisch and Lewis and, in addition, a large negative break in 1921, which accounted for a once-and-for-all decline in prices of about 40%. Figure 1 presents the evolution of the Grilli-Yang commodity price index, which has been updated to 1992.[4]

All the above papers tested the secular decline in commodity prices using a simple trend and cycle model, which assumes that prices, P_t, can be appropriately characterized by a TS representation of the form:

$$LogP_t = \beta T_t + ARMA(p,q)\epsilon_t \tag{1}$$

where T_t is a deterministic trend variable, ϵ_t is an i.i.d. random shock and parameter ß can be estimated by standard econometric procedures. The mixed autoregressive moving average (ARMA) term for the residual, used by Grilli and Yang (1988) and Sapsford, Sarkar and Singer (1992), eliminates the potential misspecification caused by higher-order serial correlation. In this model the only information required to predict the long-run evolution of the price is its mean growth rate (ps) because shocks, being totally transitory, do not affect long-run forecasts.

This methodology has been criticized after the work of Dickey and Fuller (1981) and Nelson and Plosser (1982), which suggest that the implicit assumption of stationarity is frequently rejected by economic data. If the variable is non-stationary, standard econometric procedures cannot be applied directly because the estimated parameter ps, although not biased, has a non-trivial distribution that invalidates inferences on it. Nevertheless, the first difference of prices may conform to a DS model of the form:

$$\Delta LogP_t = \beta' + ARMA(p',q')\mu_t \tag{2}$$

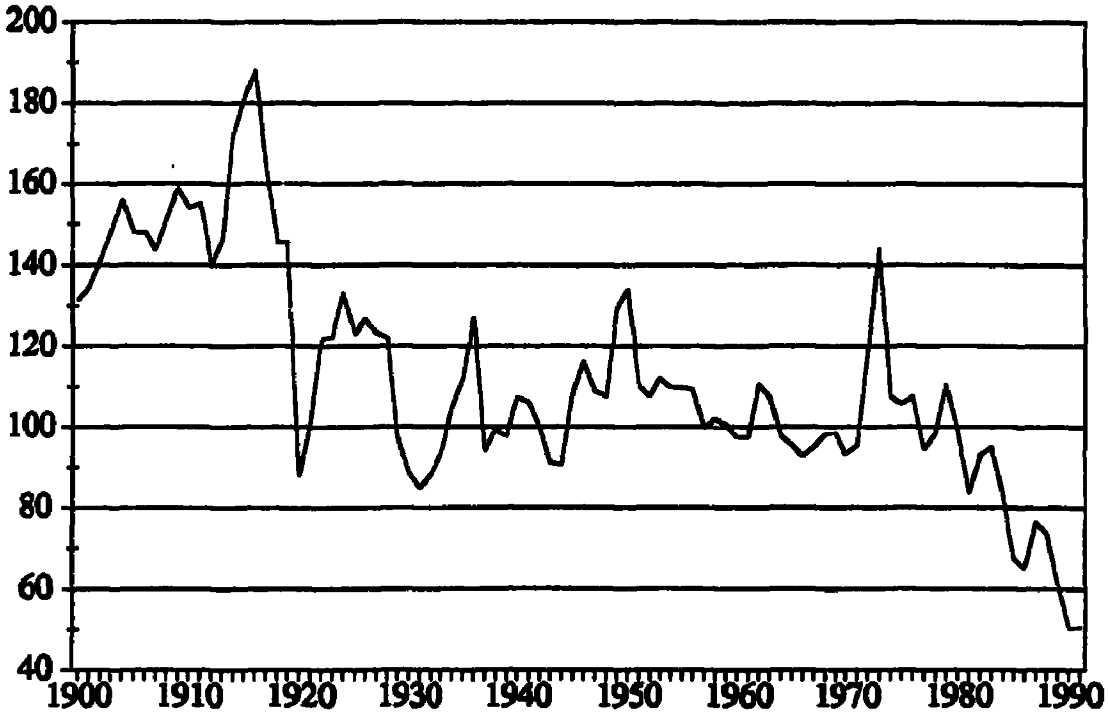
where Δ is the first-difference operator and parameter ps' is now the mean growth rate of the variable. The presence of μ_t, an i.i.d. random variable, will induce the trend of the level of commodity prices to have a stochastic behavior; hence, an important implication of this

[4] The original Grilli-Yang index comprises 24 commodities (aluminum, bananas, beef, cocoa, coffee, copper, cotton, hides, jute, lamb, lead, maize, palm oil, rice, rubber, silver, tea, timber, tin, tobacco, wheat, wool and zinc), weighted by their share in world exports in the 1977-79 period. Appendix 1 describes the data. Cuddington (1992) expanded the database to include oil and coal. In this paper, the original Grilli-Yang data was extended to include the 1987-92 period; unfortunately, data for oil and coal was not available.

FIGURE 1

Aggregate Commodity Price Index

1970 =100



specification is that, contrary to TS models, shocks may have permanent effects on the level of commodity prices, P_t.

Among the growing literature which characterizes commodity prices as integrated processes, Cuddington (1992) presents evidence that 12 of the 24 commodities which comprise the Grilli-Yang index can be modelled as non-stationary processes for the 1900-1983 period, while the rest correspond to stationary variables. Based on this classification, DS models were estimated for the former variables, while TS specification were estimated for the latter. Only five of the estimated models, all of them TS variables, presented evidence of a negative trend in the period, while the rest either were trendless or had positive trends; hence, the PSH was rejected in the majority of the cases. Regarding the persistence of shocks, gain functions obtained from the estimated ARIMA representation of the DS models suggested that a range in the persistence of shocks from 34% to 100%, where the latter corresponds to the case when commodity prices are modeled as pure random-walk processes.

When analyzing these results in depth, some elements are suggestive that the estimation may be inaccurate. In six of the 13 TS models, the fit of the estimation was low for long-horizon time-series standards (R^2 are in the order of 0.4 or 0.6), which implies that important information regarding the processes of commodity prices has been left out of the estimation. Error terms were fitted with ARMA structures with very long lags for annual data (in some cases of the 11th order), which implies that the models have problems when parameterizing the persistence of the shocks. Some implications of these results are also quite implausible from an economic point of view. In the case of the TS models it is suggested, for example, that a 1% shock to the price of coffee today will reduce the price by 0.5% in the year 2005. On the other hand, classifying copper as a pure random-walk suggests that 100% of any shock is permanent, and that the best forecast on the price at any horizon is the current price, an element which does not match easily with observed long-run cycles in the price, nor with the degree of predictability achieved by the experts (Behrman, 1987; Vial, 1992). In this case, a stabilization mechanism, such as the Chilean copper fund, would be ineffective to smooth out the instability of export revenues.

Moreover, no formal method for determining and controlling for structural breaks was included in the estimation. A level break was included for only one commodity (coffee, in

1950), apparently after visual inspection of the residuals suggested that the estimation was unstable in that year.[5] Breaks in the trend of the series were also disregarded. The next section undertakes a formal analysis of the role of structural breaks in reducing the power of unit-root tests, and implements a procedure to determine the existence of breaks in commodity prices.

3. Testing for Structural Breaks in Commodity Prices

As mentioned in the introduction, an important drawback of the standard methodology utilized to assess the long-run trends of commodity prices is its reliance on parametric unit-root tests to determine the existence of non-stationary components in the series. These tests present two types of problems: first, they usually have low power against close competitors in finite samples (Rudebusch, 1993) and, second, they are particularly sensitive to the presence of structural breaks (Hendry and Neale, 1991). The lack of power, even in the absence of breaks, arises from the fact that "any test whether a continuous parameter θ is equal to some θ_0 has arbitrarily low power against alternatives θ-ϵ in finite samples" (Cochrane, 1991). In other words, when confronted with models which are close to the unit-root case, say an AR(1) model with an autoregressive parameter of 0.96, the test will tend to reject the null hypothesis of a unit root in fewer cases than when the autoregressive parameter is 0.85. In addition, structural breaks further reduce the power of the tests: as noted by Perron (1989), a stationary model with a level shift can easily be taken for an integrated process. In fact, Perron shows that several economic series usually regarded as containing a unit root (eg, GDP, interest rates, etc) actually could correspond to stationary models with shifts in the level, the trend or both.[6]

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[5] Cuddington (1992) also introduced a break in the case of oil in 1974, a commodity not included in our study.

[6] Ardeni and Wright (1992) and Reinhart and Wickham (1994) find evidence of a secular decline in aggregate indices of commodity prices, using Harvey's structural decomposition technique, which does not need stationarity assumptions. Their results, however, depend crucially on the absence of structural breaks, as acknowledged by the authors.

Table 1
Empirical Power of the Dickey-Fuller Test with and without Structural Breaks (Percent)

AR(1) Model*

Sample Size	$\rho = 0.85$									$\rho = 0.96$								
	No Breaks			Level Break			Trend Break			No Breaks			Level Break			Trend Break		
	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%
64	8.1	30.0	48.2	0.7	5.1	10.4	7.8	28.9	45.0	1.3	7.2	13.7	1.9	4.6	9.2	1.4	0.8	1.2
128	46.6	85.0	95.1	2.6	18.0	35.8	18.9	47.4	68.3	2.9	13.3	24.4	1.2	7.0	15.6	2.9	1.7	1.4
256	99.1	100.0	100.0	32.2	81.7	95.7	12.9	49.2	73.5	11.0	36.9	54.3	8.3	20.2	33.8	4.0	13.7	27.5

ARIMA(1,1,1) Model

Sample Size	$\rho = 0.85$									$\rho = 0.96$								
	No Breaks			Level Break			Trend Break			No Breaks			Level Break			Trend Break		
	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%
64	3.0	13.3	26.1	1.5	7.1	14.9	2.5	10.9	17.7	0.7	6.5	13.1	0.9	4.2	9.5	0.8	4.8	10.8
128	5.5	18.9	29.9	2.8	14.9	26.4	4.2	16.9	28.8	1.7	7.5	15.1	2.0	6.3	12.4	2.0	7.9	15.5
256	10.4	28.5	42.0	10.4	27.0	39.0	8.1	19.4	30.3	2.7	10.6	20.3	1.8	9.1	19.1	1.7	11.6	21.2

Note: (*) The AR(1) model is specified as $X_t=\rho X_{t-1}+\mu_t$, where $\mu_t \sim N(0,1)$; the ARIMA(1,1,1) model is specified as $X_t=Y_t+Z_t$, where $Z_t=Z_{t-1}+\nu_t$ and $\nu \sim N(0,½)$ and where $Y_t=\rho Y_{t-1}+\mu_t$ and $\mu_t \sim N(0,1)$. Each set of rows corresponds to an independent and separate experiment based on 10,000 replications. Breaks of size $2\sigma_\mu$ for the level and 5% for the trend were introduced at the midpoint of each sample.

To address these issues in our case, we proceed in two stages. First, Montecarlo simulations are used to quantify the power of the most-used parametric unit-root test, the Dickey-Fuller (DF) test, in the presence of level and trend breaks. The design of the Montecarlo experiment follows that of Hendry and Neale (1991), but it has been adjusted to fit the characteristics of the commodity price data. Second, we test for the existence of breaks in the Grilli-Yang commodity prices database, using a sequential test developed by Zivot and Andrews (1992).

Table 1 presents the results of the Montecarlo simulation of the power of the DF test for different sample sizes (ranging from 64 to 256 observations), competing alternatives (AR(1) and ARIMA(1,1,1) models) and types of structural breaks (changes in the level or the trend of the series).[7] Considering first the no-breaks case, it can be seen that the power of the test when the true data generating process is a stationary AR(1) model, is low in samples of less than 128 observations, even if the specification is not close to the unit-root process (when $\rho=0.85$ the DF test correctly rejected the null hypothesis of random-walk only in 48.2% of the cases, at 10% significance levels). As expected, the power of the test declines as the model approaches the unit-root specification; when the autoregressive parameter ρ increases to 0.96, the power of the test is less than 25% for 128 observations. The more complicated ARIMA(1,1,1) model, suggested by Summers (1986) to capture slow mean-reverting processes, proves the DF test to have very low power, even for large samples of 256 observations.

To analyze the role of structural breaks we introduce two types of breaks --a level shift equivalent to 2 times the size of the standard deviation of innovations and a trend break of size 5%-- located at the mid-point of each sample size.[8] The results illustrate the significant reduction in the power of the test, in particular when the sample has 128 or fewer observations. In the

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[7] We used the pseudo random-generator routine of GAUSS to create 10,000 artificial series of two specific processes: a stationary AR(1) model with homoskedastic innovations, and an ARIMA(1,1,1) model, formed by adding to the former a nonstationary (random-walk) component. The standard deviation of the innovations in the latter was chosen to be substantially smaller than that of the stationary component (25%). We apply a standard DF test to the simulated series and use McKinnon critical values to compute the number of cases in which the test (correctly) rejects the null hypothesis of a random walk. A detailed description of the methodology and the complete set of results is presented in León and Soto (1994a).

[8] We also tested for level and trend breaks located at 25% and 75% of each sample size, but results are affected only marginally. In addition, we tested bigger size breaks which, as expected, further lowered the power of the test.

AR(1) case, the power of the test is reduced markedly in the presence of level and trend breaks, even in the specification which is far from the unit-root model (ρ=0.85): for 128 observations, the power reduces from around 95% (at 10% significance levels) to 35.8% with the level break and 68.3% with the trend shift. A less marked but still important reduction in power is observed for the specification closer to the unit root model (ρ=0.96). In the ARIMA(1,1 1) model, however, the presence of the breaks does not alter substantially the power of the test; this is not surprising since the DF test interprets the level break as a rather extreme realization of a permanent shock.

To assess the presence of structural breaks in commodity prices, we use a test that endogenizes the dating of the breaks by estimating recursively a normalized version of Perron's (1989) additive-outlier model, starting from the beginning of the period and moving period-by-period towards the end of the sample (normalization is required to control for the changing size of the sample). Perron's (1989) test cannot be used directly in this context, because in his procedure breaks are introduced by the econometrician with a-priori information only, i.e., the dating of the break is not determined by any formal procedure, apart from the visual inspection of the data. The test used in this paper allows the position of the break to be endogenously determined at the point at which the null hypothesis of a unit root with constant parameters is more easily rejected against the competing alternative TS representation with shifts in either the level or the trend of the series. The appropriate asymptotic distribution of the statistic, obtained by Zivot and Andrews (1992) using Montecarlo simulations, provides critical values well below the full-sample DF critical values. When assessing the power of the test against stationary models the recursive test has comparable power to that of the standard DF test in the absence of breaks. Its main limitation, however, is that it tests for only one break; it is, nevertheless, preferred over a similar test developed by Banerjee et al (1992) because the number of lags is endogenously determined at each recursion, avoiding potential misspecification problems.

The recursive test was applied only to the 12 commodity prices potentially subject to the problem of misclassification when using the full-sample DF tests, i.e., those cases in which a

TS model with break might have been erroneously taken to be a non-stationary process.[9] Column 1 in Table 2, which replicates Cuddington's (1992) results, shows that all variables would have been classified as DS models by full-sample DF tests; recursive tests, however, reject the null hypothesis of non-stationarity in 8 of the 12 cases, suggesting that a stationary specification with trend and/or level shifts would be more appropriate. Some of the potential breaks consist of trend shifts located around 1960, which might explain the dissimilar parameterizations of aggregate commodity price indices observed when the post World War II period is included in the sample. Interestingly, no evidence of breaks is observed for 1921, the year which has usually been designated as a break in the PSH literature. This gives some support to Cuddington and Wei's (1990) observation that the 1921 break observed in the Grilli-Yang index might be a result of aggregation problems.

These endogenous-break unit-root tests sustain the notion that the presence of structural breaks may have misled the selection of commodity price models in previous studies, and that the rejection of the PSH may be in effect due to model misspecification. An associated effect of the presence of breaks is that, if the DS/TS classification is incorrect. then the estimated measures of the persistence of the shocks are also dubious. In order to avoid these limitations, Cochrane's non-parametric estimator is used to quantify the persistence of shocks. The next section derives this estimator and performs a simple Montecarlo simulation to prove its robustness in the presence of breaks.[10]

[9] To determine ir an endogenous way the presence of breaks in TS variables, we use standard CUSUM tests, see appendix table 1.

[10] Another undesirable feature of pa-ametric techniques for measuring the persistence of shocks, such as the method of unobserved components suggested by Beveridge and Nelson (1981) or the ARMA approach developed by Campbell and Mankiw (1987), is that they concentrate on the short-run dynamic structure of the series, from which the long-run properties are *inferred*. Hence, if high-order negative autocorrelations offset initial positive autocorrelations both methods will fail to detect it, and results will be biased towards low-order DS representations.

Table 2
Potential Structural Breaks according to Zivot-Andrews Recursive Unit Root Test

Commodity	τ_{DF} Full Sample	min τ_{DF} Level Break	Date	min τ_{DF} Trend Break	Date
Aluminum	-2.63	-4.75*	1940	-4.76*	1940
Bananas	-2.28	-4.44	1924	-3.44	1932
Beef	-2.72	-4.52	1958	-3.12	1984
Cocoa	-1.98	-4.37	1947	-3.23	1923
Copper	-2.84	-4.71*	1953	-3.50	1922
Cotton	-1.30	-4.22	1985	-4.48*	1962
Jute	-2.67	-4.49	1973	-5.05*	1967
Rubber	-3.12	-4.88*	1917	-4.38*	1919
Silver	-1.83	-3.40	1962	-2.58	1931
Tea	-1.05	-4.26	1985	-5.04*	1963
Tobacco	-2.43	-4.88*	1917	-4.03	1935
Wool	-1.65	-4.07	1974	-5.69*	1952

Note: * rejects the null hypothesis of non-stationarity at 10% . The critical values are -3.45 for the full-sample test, and -4.58 for the level break and -4.11 for the trend break (Zivot and Andrews,1992).

4. A Non-Parametric Estimator of the Persistence of Commodity Price Shocks

The estimator used in this paper to measure the size of the non-stationary component in commodity prices does not rely on any parametric representation of the series. It exploits two features of time-series: first, any series whose first log-difference follows a stationary linear process can always be represented by a combination of random-walk and stationary processes, though the decomposition among components is entirely arbitrary (Watson, 1986). The TS model is just a limiting case of this representation, where the random-walk component is zero. Second, in any decomposition of the level of the series, however, the innovation variance of the random-walk component is the same (Cochrane, 1988). Hence, an estimator of the size of each component that relies on comparing the variance of the random-walk component to that of innovations is free from the problems affecting parametric tests. Moreover, since the response of the variable to innovations is proportional to the square root of the variance of shocks to the random-walk component we can easily measure the size of permanent shocks (Lo and McKinley, 1988).[11]

A simple way to obtain the variance ratio estimator consists of re-specifying equation 2 as:

$$\Delta P_t = (1-L)P_t = \mu + A(L)\epsilon_t = \mu + \sum_{j=0}^{\infty}\theta_j \epsilon_{t-j} \tag{3}$$

where the TS representation (equation 1) can be recovered by setting $\mu=\beta$ and A(L)=ARMA(p-1,q). From Beveridge and Nelson (1981) we know that every integrated series admits at least one decomposition of equation (3) into permanent and transitory components (Z_t and C_t, respectively), which is guaranteed to exist:

[11] A key element in this decomposition is the notion that the permanent component can be represented as a random walk. In more general specifications, however, the variance of the permanent component cannot be identified from the second moments of the original series, i.e., identification requires auxiliary assumptions (see Quah, 1991).

$$
\begin{aligned}
P_t &= Z_t + C_t \qquad\qquad\qquad \textit{where} \\
Z_t &= \mu + Z_{t-1} + \left(\sum_{j=0}^{\infty} \theta_k\right) \epsilon_t \\
-C_t &= \left(\sum_{j=1}^{\infty} \theta_k\right) \epsilon_t + \left(\sum_{j=2}^{\infty} \theta_k\right) \epsilon_{t-1} + \left(\sum_{j=3}^{\infty} \theta_k\right) \epsilon_{t-2} + \ldots
\end{aligned}
\tag{4}
$$

The permanent component is defined in this representation as the long-run forecast value of P or, in terms of present and past shocks, as the cumulative effect of disturbances. The variance of the first difference of the permanent component can be obtained easily from equation (4) as: $\sigma^2_{\Delta Zt} = |A(1)|^2 \sigma^2_{\epsilon t}$. The V-estimator is then defined as the ratio of the variance of the first differences of the non-stationary component to the variance of the first differences of the series, i.e., the variance of innovations ($\sigma^2_{\Delta Pt}$):

$$
V = \frac{\sigma^2_{\Delta Z_t}}{\sigma^2_{\Delta P_t}} \tag{5}
$$

Interpretation of V as a measure of the importance of the random-walk component is straightforward. If P_t is a strictly stationary process in levels, $\sigma^2_{\Delta zt}$ will be zero at any time because there is no permanent component, therefore, V=0. On the other extreme, if P_t is a pure random-walk process, $\sigma^2_{\Delta zt}$ will be equal to $\sigma^2_{\Delta Pt}$, because all shocks are permanent; hence, V=1. Moreover, V can be bigger than one if the series exhibits an explosive behavior.

An alternative way to express the variance ratio, used in the empirical analysis, can be obtained using the autocorrelation coefficients of the first differences of the series.[12]

[12] A formal derivation is contained in Appendix 2.

$$V^{m} = 1+2\sum_{j=1}^{m}\frac{(m-j)}{m}\rho_{\Delta P_j} \qquad where \quad \rho_{\Delta P_j} = \frac{cov(\Delta P_{t-j}\Delta P_t)}{\sigma^2_{\Delta P}} \tag{6}$$

where the term (m-j)/m adjusts for sample size. Evidently, if P_t is a pure random-walk process, then its first differences must be serially uncorrelated --so that $\rho_{\Delta Y_j} = 0$ for all $j>0$-- and V is equal to one. Values of V below one indicate the existence of negative autocorrelations of first differences, i.e., that the series presents mean reversion. The selection of the number of autocorrelations *m* in equation (6) is entirely arbitrary; nevertheless, a reduced number of autocorrelations may induce a failure to detect mean reversion, because of the presence of high positive autocorrelations at lags near to zero. Following standard procedures, the maximum *m* in the empirical section was set at one-half of the sample size.

Several characteristics of this estimator are worth noting. The specification is less restrictive than that of Beveridge and Nelson (1981), in that here it is not required that permanent and transitory components be uncorrelated. Equation (6) highlights the long-run nature of the estimator, as it uses a large number of autocorrelations of the first differences, and not only those lags near to zero as most parametric tests. In addition, since this estimator is based on the autocorrelation function of the series, it is to a great extent less sensitive to non-normal disturbances and, more important, to structural breaks, as discussed below.

Lo and McKinley (1988) and Chow and Denning (1993) report that the power of the V test against the AR(1) alternative is comparable to that of parametric tests, while it outperforms them when dealing with ARIMA and ARMA models, and also that it performs satisfactorily in cases where errors are correlated or heteroskedastic. To our knowledge, the power of the V test under structural breaks has not been studied in detail. Given the analytical intractability of the problem, we perform a Montecarlo simulation of the power of the test and present the results in table 3.[13]

Following Chow and Denning, we present the combined results for different numbers of autocorrelations (ranging from 8 to 64). It is apparent that, when confronted with the AR(1) model, the V^m statistic possesses a similar power to that of the DF in large samples and slightly less power in smaller samples (see also table 1). In the case of the ARIMA(1,1,1) model, both tests perform rather poorly when the model is closer to the random-walk specification (ρ=0.96); nevertheless, the normalized ratio of variances is more powerful when the stationary component is larger in the alternative model (ρ=0.85). Regarding the power of the V^m statistic in the presence of level and trend breaks, the ratio of variances clearly outperforms the DF test in the AR(1) specification in large and small samples, in particular when the model has ρ=0.85; in the case of the ARIMA(1,1,1) model, both tests have similar power. In the case of trend shifts, however, the V^m statistic is clearly superior in the AR(1) model, and also in the ARIMA(1,1,1) model when ρ=0.85.

In summary, this exercise suggest that though the ratio of variances is not more powerful than the DF test in the absence of structural breaks, it is clearly more robust to level and trend shifts, even if they are of a large magnitude. This is an important advantage of the non-parametric estimator over standard parametric methods when analyzing long-run behavior, particularly in the case of commodity prices which, as discussed in section 2, have experienced structural breaks during the last 90 years.

---

[13] The experiment is entirely equivalent to that reported in table 1. In the simulations, however, we use a normalized version of V, developed by Chow and Denning (1993), which is easier to test because it has a known normal distribution:

$$Z(m) = \sqrt{T}\cdot V^m\cdot\left(\frac{2(2m-1)(m-1)}{3m}\right)^{1/2} \sim N(0,1)$$

Table 3
Empirical Power of the Ratio of Variances Test with and without Structural Breaks (Percent)

AR(1) Model*

Sample Size**	$\rho = 0.85$									$\rho = 0.96$								
	No Breaks			Level Break			Trend Break			No Breaks			Level Break			Trend Break		
	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%
64	4.5	19.0	32.8	2.1	11.3	21.7	3.3	18.3	32.8	1.2	5.8	11.8	0.9	5.1	10.4	0.9	6.7	11.5
128	21.3	56.4	75.0	11.6	36.7	55.3	16.6	48.7	68.6	1.7	8.9	17.3	1.5	8.1	15.5	1.7	8.9	16.7
256	86.5	98.1	99.5	65.1	92.1	98.9	77.5	96.7	98.9	5.4	20.7	34.7	4.6	17.6	30.5	4.9	20.9	32.3

ARIMA(1,1,1) Model

Sample Size	$\rho = 0.85$									$\rho = 0.96$								
	No Breaks			Level Break			Trend Break			No Breaks			Level Break			Trend Break		
	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%	1%	5%	10%
64	2.4	10.9	21.7	1.8	8.8	17.1	2.4	13.0	19.7	1.1	5.7	11.0	0.9	5.0	10.1	0.8	5.7	11.6
128	5.5	21.3	34.9	4.4	17.9	30.9	5.6	20.4	34.5	1.5	7.1	13.9	1.3	6.8	13.5	1.7	7.1	13.8
256	18.3	45.7	63.0	15.7	41.4	58.1	15.8	41.7	58.5	2.5	2.1	22.2	2.3	11.2	20.3	2.9	11.4	20.8

Notes: (*) The AR(1) model is specified as $X_t=\rho X_{t-1}+\mu_t$, where $\mu_t \sim N(0,1)$; the ARIMA(1,1,1) model is specified as $X_t=Y_t+Z_t$, where $Z_t=Z_{t-1}+\nu_t$ and $\nu \sim N(0,½)$ and where $Y_t=\rho Y_{t-1}+\mu_t$ and $\mu_t \sim N(0,1)$. Each set of rows corresponds to an independent and separate experiment based on 10,000 replications. Breaks of size $2\sigma_s$ for the level and 5% for the trend were introduced at the midpoint of each sample. (**) For 64 observations, we compute m=8, 16, and 32 autocorrelations; for 128 and 256 observations, we compute m=16, 32, and 64 autocorrelations.

5. How Big Is the Random-Walk Component in Commodity Prices?

Figure 2 presents a graphic display of the normalized variance ratio estimated for the group of 24 commodities, where dotted lines represent 5% bands of confidence. The first conclusion is that, contrary to the evidence collected through full-sample DF tests, only five of the series (aluminum, bananas, cotton, tobacco, and wool) present the typical pattern of processes containing a large unit-root, where the long-run variance ratio approaches or exceeds 1 at long horizons. In all but one case (tobacco) the estimated V^m does not reach 1, but we cannot reject the null hypothesis of non-stationary, because the upper 5% confidence band exceeds 1 in the long-run. On the contrary, the other 19 variables present clear evidence of important mean-reverting processes, which support the notion that, in order to capture long-run trends with only 93 observations, it would be better to use TS models.

A useful way to summarize this evidence is to classify the dynamics of commodity prices according to the long-run ratio of variances, i.e., the value at which the statistic reaches a stationary plateau. Among the TS variables, thirteen commodity prices present very low levels of persistence --less than 35% of the total variance of innovations is caused by the random-walk component--, while other six display moderate persistence (between 35% and 50%). A comparison of these results to that of parametric tests reveals important discrepancies. Only 5 of the 12 commodity prices originally classified as DS processes by DF tests coincide with those suggested by the non-parametric criteria (aluminum, bananas, cotton, tobacco, and wool); while for the remaining 7 cases (beef, cocoa, copper, jute, rubber, silver, and tea) the evidence obtained by using the V^m estimator suggests the use of TS models. In the case of 4 commodity prices (copper, jute, rubber and tea), the evidence from the non-parametric estimator supports the Zivot-Andrews DF results, namely that the DS specification can be rejected against a TS model with structural breaks. On the other hand, 3 commodity prices classified by full-sample and Zivot-Andrews DF tests as DS models (beef, cocoa, and silver) present evidence of important mean-reversion processes, which would favor a TS representation.

The inability of parametric tests to capture slow mean-reversion processes is apparent when analyzing the speed of convergence of shocks to their long-run persistence, illustrated by the slope of the V^m function. It is apparent that in the three cases in which parametric tests give

Figure 2
Estimated Variance Ratios
for Commodity Prices, 1900-1992

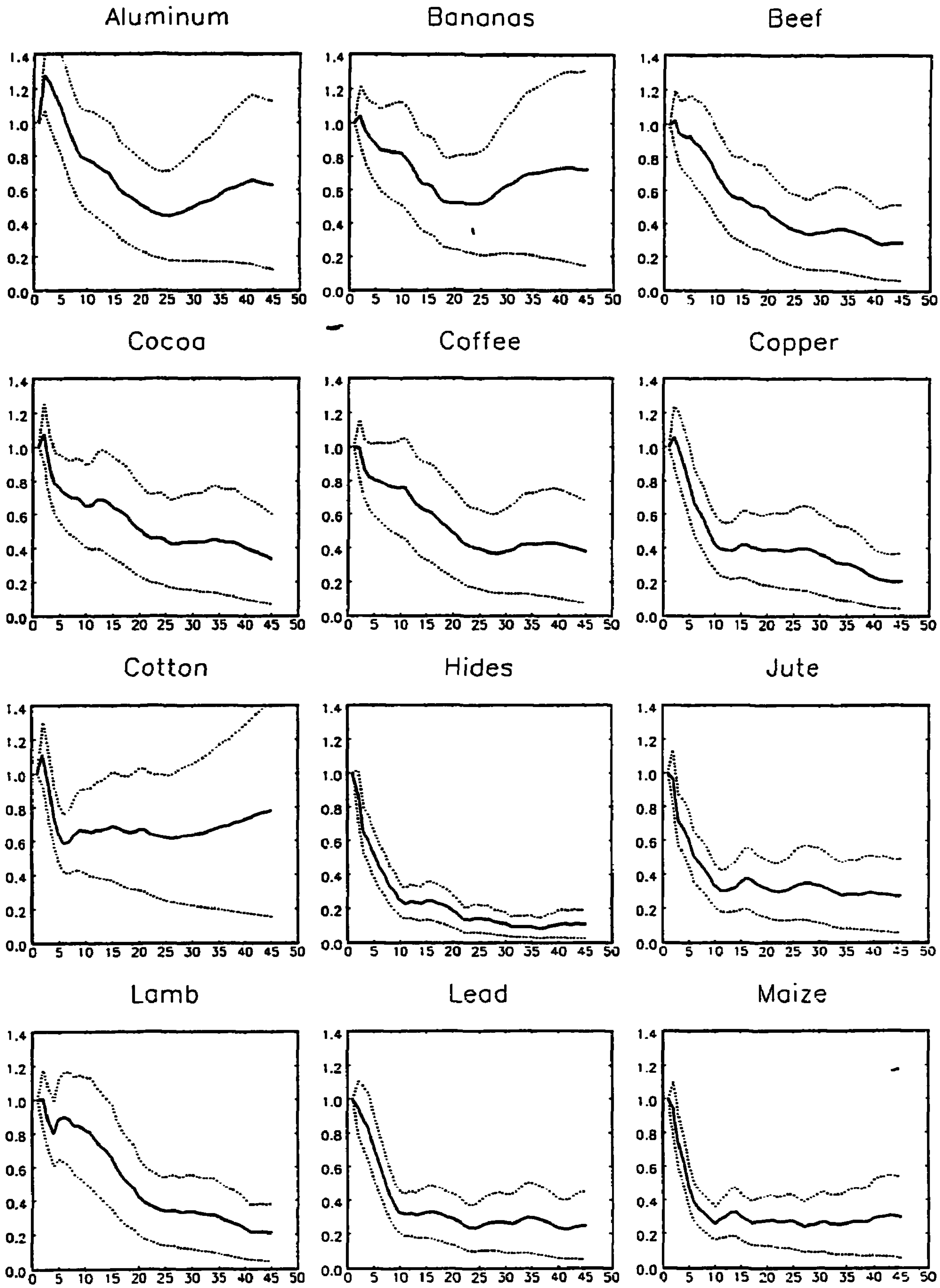
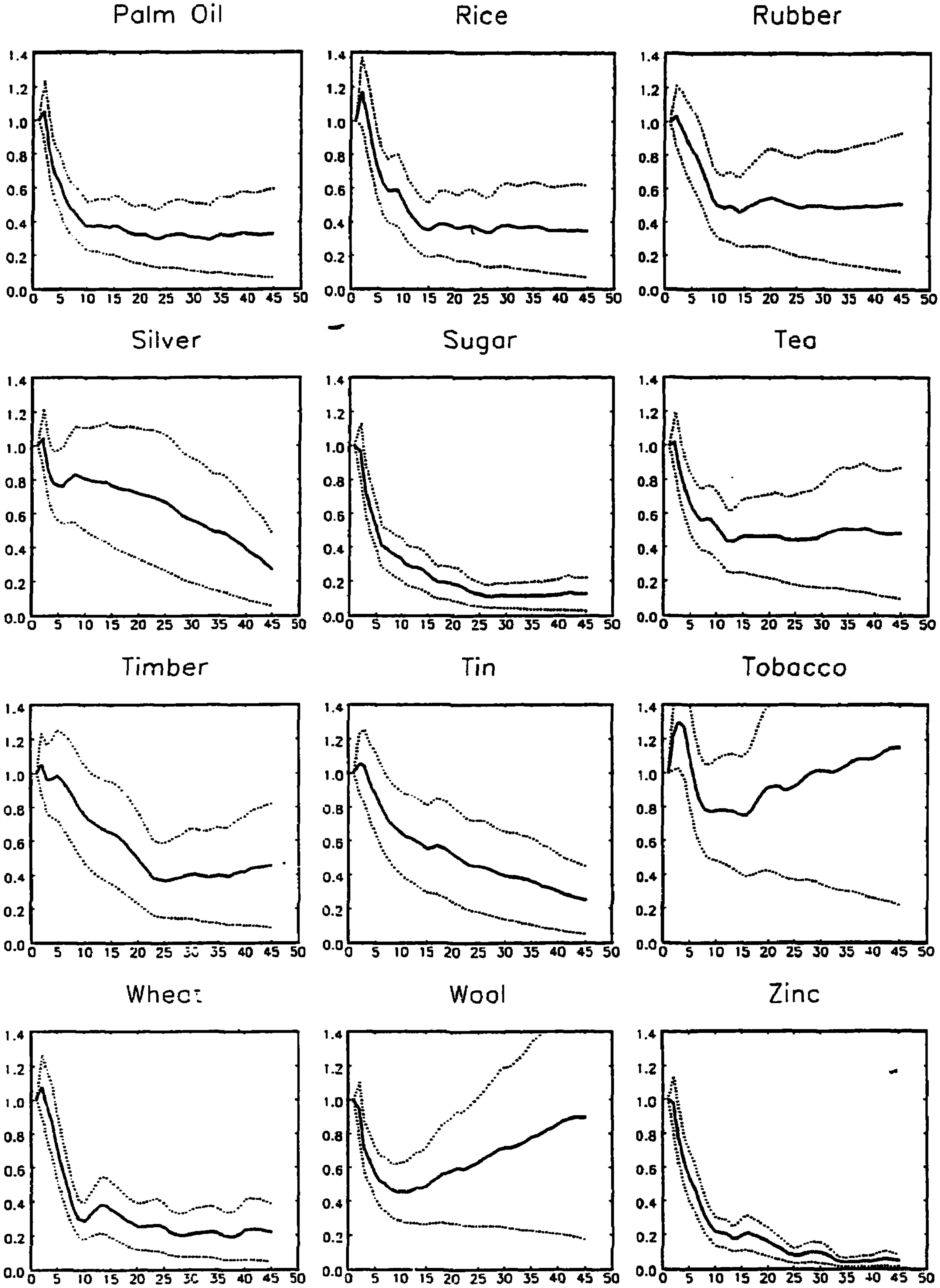


Figure 2 (cont.)



opposite results to that of the V^m estimator (beef, cocoa, and silver), shocks present slow convergence patterns. If the mean reverting process is slow, the DF test tends not to reject the null hypothesis even if persistence is low, because the high-order negative autocorrelations that drive the variable back to its trend are not taken into account. On the other hand, there are cases in which the mean reverting process is fast and could be captured by the DF test, but the procedure fails because the random-walk component is small but important compared to the stationary component (rubber and tea). Finally, when the mean reverting process is fast, i.e. when shocks dissipate quickly, the DF tends to correctly capture the process because high-order autocorrelations are not important; in all cases in which DF tests rejected the null hypothesis of a unit root, the slope of the V schedule is steep in the first periods (eg, maize, sugar, wheat, zinc).

This evidence, in addition to that provided by structural break unit-root tests in section 2, suggests that previous modeling of commodity prices in small samples has been inappropriate in most cases. Table 4 presents new estimates of long-run trends in commodity prices based on this non-parametric classification, and taking into account the structural breaks found in the series according to table 2 and appendix table 1. The results show an important difference from previous results in the literature. Of the nineteen TS models, 11 present a negative and significant trend, in the range of -0.2% to -1.2% per year, five are trendless processes and three variables have a clear positive trend. In the case of DS models, one commodity price presents no indication of a stochastic trend (aluminum). The other DS models (bananas, cotton, tobacco, and wool) do not present a stable trend process; in all cases there is no drift in the pre-break period, while a negative drift is found in the post-break period. In summary, 15 of the 24 commodity prices present a negative trend for all or most of the 1900-92 period.

It is important to highlight the different economic implication of these results when compared to those of Cuddington (1992). First, using the V^m as a measure of persistence reduces much of the risk of confusing permanent innovations with structural breaks, hence diminishing the bias towards overestimation. For example, Cuddington's measures of the persistence of shocks of copper and rubber prices is 100%; if structural breaks are included, TS models would have to be estimated, which in turn would imply zero persistence for both. The non-parametric estimator suggests that the level of persistence is 22% for copper and 52% for rubber. These

Table 4
Estimated Long-Run Trends of Commodity Prices, 1900-1992

Commodity	Model	Constant	Trend	Level Dummy	Trend Dummy	Error Process	R^2	Q(12)
Aluminum	DS	-0.015 (-0.97)				$e_t=(1+0.37L-0.30L^2-0.25L^3)\mu_t$ (3.36) (-2.78) (-2.31)	0.18	3.14
Bananas	DS	-0.013 (-1.31)			-0.034 (-2.00)	$(1+0.19L^3)\,e_t=\mu_t$ (1.78)	0.07	9.51
Beef	TS	-9.11 (-1.03)	0.004 (1.00)	0.89 (4.60)		$(1-0.75L)\,e_t=\mu_t$ (-10.49)	0.91	9.85
Cocoa	TS	-1.10 (-4.19)	0.001 (0.25)			$(1-0.78L)e_t=(1-0.33)\mu_t$ (-13.89) (-3.21)	0.76	11.2
Coffee	TS	11.05 (12.88)	-0.005 (-0.91)	0.45 (2.00)		$(1-0.81L)\,e_t=\mu_t$ (-10.96)	0.73	8.42
Copper	TS	24.4 (9.56)	-0.012 (-9.52)	0.67 (9.49)		$e_t=(1+0.59)\mu_t$ (7.10)	0.66	22.3
Cotton	DS	0.002 (0.19)			-0.045 (-2.57)	$(1+0.71L+0.31L^2)$ $e_t=(1+0.80L)\mu_t$ (4.49) (2.32) (7.48)	0.09	12.69
Hides	TS	20.42 (4.35)	-0.010 (-4.34)			$(1-0.61L)\,e_t=\mu_t$ (-7.37)	0.65	9.34
Jute	TS	-3.11 (-0.59)	0.002 (0.64)		-0.0002 (-3.05)	$(1-0.56L)\,e_t=(1-0.45L)\mu_t$ (-6.99) (-4.45)	0.65	13.99
Lamb	TS	-34.82 (-4.12)	0.017 (4.02)			$(1-0.91L+0.14L^2)e_t=\mu_t$ (-13.76) (2.17)	0.87	18.29
Lead	TS	19.51 (3.21)	-0.010 (-3.25)	0.93 (2.77)		$(1-0.70L)\,e_t=\mu_t$ (-8.88)	0.65	5.04
Maize	TS	15.31 (1.92)	-0.008 (-1.89)	-0.48 (-2.45)		$(1-0.78L)\,e_t=\mu_t$ (-11.33)	0.71	15.28
Palm oil	TS	11.26 (3.31)	-0.006 (-3.59)	-0.86 (-5.99)		$(1-0.78L-0.24L^2)\,e_t=\mu_t$ (-7.43) (2.37)	0.82	4.06
Rice	TS	23.75 (5.89)	-0.012 (-5.85)	0.22 (1.81)		$(1-0.61L)\,e_t=(1+0.51L)\mu_t$ (-8.82) (5.33)	0.82	6.69
Rubber	TS	36.99 (4.68)	-0.018 (-4.64)	-0.89 (-3.98)		$(1-0.89L+0.22L^2)\,e_t=\mu_t$ (-8.51) (2.01)	0.92	10.59
Silver	TS	-0.81 (-010)	0.0002 (0.005)			$(1-0.34L-0.49L^2)\,e_t=(1+0.70L)\mu_t$ (-2.62) (-3.75) (7.46)	0.81	7.79
Sugar	TS	10.21 (1.95)	-0.005 (-1.93)	-0.36 (-2.21)		$(1-0.31L)\,e_t=(1+0.48L)\mu_t$ (-3.34) (4.65)	0.60	11.23
Tea	TS	0.32 (0.05)	-0.001 (-0.04)		-0.001 (-1.94)	$(1-0.51L-0.21L^2)\,e_t=(1+0.78L)\mu_t$ (-4.15) (-1.75) (10.1)	0.77	10.27
Timber	TS	-17.47 (-3.14)	0.008 (3.02)	-0.31 (-2.49)		$(1-0.81L)\,e_t=\mu_t$ (-12.09)	0.86	7.78
Tin	TS	-20.11 (-4.23)	0.010 (3.96)	-0.78 (-5.27)		$(1-0.90L+0.18L^2)\,e_t=\mu_t$ (-9.00) (1.83)	0.82	2.94
Tobacco	DS	0.005 (0.51)		-0.31 (-3.33)		$(1-0.22L+0.28L^2)\,e_t=\mu_t$ (-2.12) (2.59)	0.25	9.71
Wheat	TS	17.96 (5.85)	-0.009 (-5.74)			$(1-0.88L+0.30L^2)\,e_t=\mu_t$ (-8.61) (2.82)	0.77	8.59
Wool	DS	0.008 (0.49)			-0.056 (-2.36)	$(1+0.46L+0.27L^2)\,e_t=\mu_t$ (4.19) (2.45)	0.20	2.15
Zinc	TS	0.08 (0.03)	0.0004 (0.34)	-1.04 (-7.49)		$(1-0.78L+0.28L^2)\,e_t=\mu_t$ (-7.41) (2.66)	0.61	8.27

Notes: t-statistics in parenthesis; the critical value for the Q test is 33.9 at 95% confidence; dummies corresponds to those in table 3 and appendix table A.1.

opposite characterizations are of importance for policy purposes; when facing an adverse price shock, a policy maker would surely react very differently if the shock is perceived as permanent than if it is expected to last only 3 to 5 years. Second, we also avoid most of the problems associated with the ARMA parameterization of the residuals implicit in gain functions, which usually disregard high-order effects. For example, the ARIMA representations of two DS models suggest persistence levels of around 84% for cotton and 89% for tobacco, while non-parametric estimates are only 65% and 76%, respectively.[14]

Regarding the PSH, the main conclusion of the analysis of long-run trends in commodity prices is that several variables contain a negative trend, although this is by no means a stylized fact. The average trend for all commodities with deteriorating trends is about -1.5% per year, comparable to that obtained with aggregate data. However, the average trend for variables with improving terms of trade is 0.9% per year, so that the resulting weighted average trend for the group of commodities is smaller than previous aggregate estimations (-0.2% per year or a cumulative 16.1% in the 93-year period).

An important caveat in the interpretation of these results is that finding negative or positive trends in some commodity prices does not automatically validate or reject the PSH because, as an indicator of LDCs trade stance, commodity prices are not an optimal proxy. First, it is likely that in a long period of time, countries might have changed the composition of their exports and imports in response to changes in relative prices; hence, a decline in the price of an individual commodity is by no means an indication that the situation has become more adverse for the country exporting that commodity. Second, LDCs do not export only primary commodities nor import only manufactures and, in fact, most countries have been diversifying exports over the past two decades. Third, welfare implications are hard to assess; if the costs

[14] It is tempting to compare the persistence as implied by the V^m measure and that arising from the gain function of the DS models; it would, however, be incorrect to undertake such comparison for the reasons explained in section 2. The gain function does not capture mean reversions arising from high order negative autocorrelation and, hence, it tends to overestimate the true persistence.

of producing primary goods decline more than those of manufactures, a primary producer country may be better off even if a decline in prices is observed.[15]

6. Concluding Remarks

This paper explores the long-term dynamics of real commodity prices using a combination of non-parametric and parametric techniques, designed to overcome several limitations of stochastic-trend models. The methodology emphasizes two aspects of the time series analysis usually overlooked in the terms of trade literature which, nevertheless, prove to be crucial when estimating long-run trends. First, we explicitly test the existence of structural breaks in each commodity; moreover, we determine the dating of breaks by using a formal test rather than by visual inspection of the data, thus avoiding implicit biases in previous papers. This is an important consideration when working with long horizons, as changes in the economic structure are likely to affect commodity price structures. Second, we show that parametric approaches give a misleading picture of long-run trends in prices, because by concentrating on short-run dynamics they lose track of mean-reversion processes.

Following Cochrane (1988) we obtain a non-parametric measure of the mean-reverting process and the speed at which shocks dissipate on time, a feature not sufficiently explored in previous papers. The estimated level of persistence is then used to select an appropriate representation of the long-run trend of prices. A taxonomy of the dynamic responses to shocks, in terms of persistence and their speed of convergence, provides a rich framework for future analysis of the determinants of commodity prices.

The empirical results alter significantly the policy implications arising from trend-and-cycle and stochastic-trend models. We show that in several cases shocks to commodity prices are far less persistent than what was suggested by previous literature (eg, Cuddington, 1992;

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[15] Some of these problems can be avoided by using directly the observed trade-weighted terms of trade of the country. León and Soto (1994b) explore these issues for 16 Latin American countries in the 1928-93 period and find that, in general, there is little evidence of a systematic deterioration in the terms of trade.

Reinhart and Wickham, 1994); hence, there may be more room for stabilization and/or price support mechanisms, as the consumption smoothing gains can offset eventual efficiency costs. By the same token, we refute the notion of shocks as purely transitory and, thus, that authorities can fully offset their adverse effects by choosing an appropriate policy mix. In addition, it is important to notice that for some goods -notably cocoa and coffee- there is evidence of a rather slow speed of dissipation of transitory shocks, which in turn suggests that commodity funds may encounter financial difficulties in the presence of extreme adverse shocks.

Finally, with regards to the Prebisch-Singer hypothesis results are mixed: though the majority of cases exhibit a negative long-run trend there are cases of zero or positive trends, rendering this negative trend a common, though not universal phenomenon. For the aggregate group of goods we estimate a cumulative price decline of around 16% for the 1900-92 period, roughly one third of the estimates obtained in previous studies (Grilli and Yang, 1988; Ardeni and Wright, 1992).

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APPENDIX 1
DATA DEFINITIONS AND SOURCES

Annual data on commodity prices for the period 1900-1988 were obtained from the Grilli and Yang database and extended to 1992 using the IMF commodity price database. The 24 commodities included in the index are: aluminum, bananas, beef, coffee, cocoa, copper, cotton, hides, jute, lamb, lead, maize, palm oil, rice, rubber, silver, sugar, tea, timber, tin, tobacco, wheat, wool and zinc. The aggregate index is a weighted average of nominal prices, where the weights were determined according to the 1977-1979 trade volumes. The price index of manufactures was obtained from the same sources and corresponds to the manufactures unit value, calculated by United Nations (MUV). An alternative deflator (the Manufacturing Price Index of the US) did not alter the results in a significant manner.

APPENDIX 2
DERIVATION OF THE VARIANCE RATIO STATISTIC IN TERMS OF THE AUTOCORRELATION FUNCTION

This appendix follows Cochrane's (1988) derivation of the V^m estimator from the autocorrelation function of the series. Start with equation (3)·

$$(1-L)P_t = \beta' + A(L)\mu_t = \beta' + \sum_{j=0}^{\infty} \theta_j \mu_{t-j}$$

using $(1\text{-}L^k)(1\text{-}L)^{-1} = (1+L^2+L^3+\ldots+L^{k-1})$ we get:

$$P_t - P_{t-k} = k\beta' + \sum_{j=0}^{k-1}(\sum_{l=0}^{j}\theta_l)\mu_{t-j} + \sum_{j=k}^{\infty}(\sum_{l=j-k+1}^{j}\theta_l)\mu_{t-j}$$

taking its variance:

$$\sigma_k^2 = k^{-1}\, var(P_t - P_{t-k}) = k^{-1}\,[\sum_{j=0}^{k-1}(\sum_{l=0}^{j}\theta_l)^2 + \sum_{j=k}^{\infty}(\sum_{l=j-k+1}^{j}\theta_l)^2]\sigma_\mu^2$$

expressing the variance as a difference equation:

$$k\sigma_k^2 - (k-1)\sigma_{k-1}^2 = [\sum_{j=0}^{\infty}(\theta_j^2 + 2\theta_j\sum_{l=1}^{k-1}\theta_{j+l})]\sigma_\mu^2$$

and letting:

$$\sigma_l^2 = (\sum_{j=0}^{\infty}\theta_j^2)\sigma_\mu^2$$

we get

$$1 + 2\sum_{j=1}^{k-1} \rho_j = \frac{k\sigma_k^2 - (k-1)\sigma_{k-1}^2}{\sigma_k^2}$$

where ρ_j is the j-th autocorrelation of $(1-L)Y_t$. Therefore:

$$\frac{\sigma_k^2}{\sigma_l^2} = k^{-1}[1 + (1+2\rho_1) + (1+2\rho_1+2\rho_2) \ldots] = 1 + 2\sum_{j=1}^{k-1} \frac{k-j}{k}\rho_j$$

Appendix Table 1
Cusum Break Tests for TS models

Commodity Price	Type of Break	Date
Coffee	Level	1945
Maize	Level	1920
Palm oil	Level	1985
Rice	Level	1920
Sugar	Level	1922
Timber	Level	1985
Tin	Level	1985

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